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ABSTRACT

This paper examines the following issue: Is the process of educational attainment the same for both whites and Mexican Americans, or does it differ? Two possible explanations for why mean differences exist in educational outcomes for whites and Mexican Americans are that either the process of educational attainment varies between the two groups, or if the process is invariant, one group starts with social advantages not shared by the other group. Data for the study were drawn from the National Longitudinal Study of the High School Class of 1972. The basic model of educational attainment used in the analysis considers education to be a function of father's occupational status, father's education, mother's education, number of siblings, sex, ability, academic preparation, and college plans. An examination of these variables indicates that there are more similarities between the two groups than there are differences. There are some differences, but it cannot be said that they consistently favor or disfavor either group. For example, for neither group is the effect of father's occupation statistically significant. For Mexican Americans the influence of father's education is nearly twice that exhibited among whites, but the influence of mother's education among Mexican Americans is negligible. The overall impression of the study is that the process of educational attainment may differ between the two groups, but not a degree worthy of much notice. (Author/RM)

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MEASUREMENT ERROR MODELS OF EDUCATIONAL ATTAINMENT FOR WHITES AND MEXICAN-AMERICANS *

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MEASUREMENT ERROR MODELS OF EDUCATIONAL
ATTAINMENT FOR WHITES AND MEXICAN-AMERICANS

ABSTRACT

There are fundamentally two possible explanations for why mean differences exist in educational outcomes for whites and Mexican-Americans: either the process of educational attainment varies between the two groups, or if the process is invariant, one group starts with social advantages not shared by the other group. Using data from the National Longitudinal Study of the High School Class of 1972, and a statistical procedure which controls for differential measurement errors between groups, we found statistically significant interactions across groups, but the differences were not large. Differences in years of schooling between whites and Mexican-Americans are mostly a function of differential levels of social background.

Measurement Error Models of Educational
Attainment for Whites and Mexican-Americans

In 1970, the median number of school years completed for white Americans was 12.2; in comparison, people of Spanish surname had completed only 8.6 years. These differences result from differential access to education by white Americans and Mexican-Americans, and there are two ways to explain such differences. On the one hand, the social process by which people in this country translate their social and biological characteristics into years of schooling may be the same for both groups, but the mean level of these characteristics may differ. In other words, some people are born with social advantages while others are not; thus the outcome may be different even though the process is the same. On the other hand, the process by which people acquire years of schooling may differ between the two groups. Regardless of mean levels of background variables, it may be that Mexican-Americans cannot translate their human capital into years of schooling as readily as can whites. Furthermore, if differences exist in the process of educational attainment between whites and Mexican-Americans, the differences may be in how the groups translate their social background into years of schooling, or in how the groups are differentially processed through the educational system. These are the main issues addressed in this paper. Is the process of educational attainment the same for both whites and Mexican-Americans, or does it differ?

BACKGROUND OF THE STUDY

Among whites, the process of status attainment in the United States is by now well understood. The seminal work of Blau and Duncan (1967) showed that about one-third of the variation in occupational status could be explained by a small set of predictor variables, the most important effect coming from respondent's educational attainment. The decade since Blau and Duncan's (1967) report has seen a large number of similar analyses extending and modifying the basic model of the process of achievement. The most important of these include Duncan, Featherman and Duncan (1972), Jencks, et al. (1972), Sewell and Hauser (1975), Hauser and Featherman (1977), and Featherman and Hauser (1978).

While a great deal of attention has been given to occupational achievement, educational attainment has also been a major focus of inquiry. Education is not only an important event in the process of occupational placement, it is equally one of the more important outcomes of the process of achievement. Major inquiries into the process of educational attainment, such as Hauser (1971), Sewell and Hauser (1975), Sewell, Hauser and Featherman (1976), and others, reveal that nearly half of the variation in educational attainment can be explained by such variables as parental education, father's occupation, respondent's intelligence, grades, curricular placement, educational and occupational aspirations, and the like.

Analyses such as those referenced above have been based for the most part on samples of white males. A few studies have addressed differences between whites and blacks (e.g., Duncan, 1969; Portes and Wilson, 1976). Even fewer studies have addressed differences in the

achievement process between whites and Mexican-Americans. Several such studies have investigated the educational aspirations of Mexican-Americans; not unexpectedly, the whites in these samples were in general more likely to express aspirations for more schooling than were Mexican-Americans, and the effects of various predictor variables were different between the two groups (Heller, 1964; TenHouten, et al., 1971; Ovando, 1978).

In terms of educational attainment, Cantu' (1975) studied a small sample of subjects from Mercedes, Texas. Among social background variables, parents' income and education were found to be important determinants of educational attainment; among social-psychological variables, parental encouragement to attend college and self-motivation were important.

Featherman and Hauser's (1978) more comprehensive study of socio-economic achievement among U. S. men reveal important differences in the process of educational attainment between Mexican-Americans and blacks. Factors such as father's education and farm background were found to be more important determinants of educational attainment for Mexican-Americans than for blacks because the mean levels of these variables were less for Mexican-Americans than for blacks. Featherman and Hauser (1978: 466) concluded that there were greater educational opportunities for blacks in the U. S. than for Mexican-Americans. In other words, their status origins were more of a handicap for Mexican-Americans than for blacks.

Featherman and Hauser's (1978) results were based on a large sample of U. S. adult males. There are several unresolved issues which make the following analysis a useful addition to this line of research. One of these is whether differences in the process of educational attainment exist among recent cohorts of Americans. A second unresolved issue is how the process of educational attainment differs between white Americans and

Mexican-Americans for a representative national sample of young adults. And the third issue is the methodological one of how to compare structural coefficients between two groups when they exhibit differential levels of measurement error. If the levels of measurement error differ between two groups, part of the observed differences in structural coefficients would be due, not to real differences in the process itself, but to differences in reliabilities of measurement. Bielby, Hauser and Featherman (1977) found differences in reliability of measurement between blacks and whites, and concluded that these measurement differences exaggerated social differences in measures of returns to schooling. In addition, Wolfle (1979) showed that Hispanics also report social data with inherent measurement errors, and ignoring them would lead to estimates even more biased than among either whites or blacks.

DATA BASE

Data for this study are drawn from the National Longitudinal Study of the High School Class of 1972 (see Levinsohn, et al, 1978). The NLS was designed to provide data on the development of the educational, vocational, and personal aspects of the lives of adolescents as they make the transition from high school to the adult world. The initial sampling frame consisted of 1200 schools, two from each of 600 strata; 18 students were selected from each school, for a total sample of 21,600 high school seniors. Several revisions of the initial design were made necessary by various logistical problems. The most important of these was the failure of the initial survey to collect data on nearly 6000 students. As a result, some important base-year responses are missing for the 6000 students, and the following analysis is

based on the 16,683 remaining students. The sample was further restricted to subjects whose racial-ethnic identification was either white or Mexican-American. In the latter case it was decided not to aggregate groups of Cubans, Puerto Ricans, and others of Spanish descent into a single group of Hispanics due to the diverse nature of their backgrounds and cultural heritage. Rather, we included only those who identified themselves as Mexican-American. As with most other analyses of the process of socio-economic achievement, we used pairwise present correlations to estimate the parameters of our model; the average number of whites in the analysis was 11,743; the average number of Mexican-Americans was 493.

SPECIFICATION OF THE MODEL

The basic model of educational attainment used in our analysis considers education to be a function of father's occupational status, father's education, mother's education, number of siblings, sex, ability, academic preparation, and college plans. The model is shown diagramatically in Figure 1; the theoretical variables of interest are shown within the ellipses. In the diagram, straight, unidirectional arrows represent hypothesized causal relationships; the arrows point toward the dependent variable. For the sake of diagramatic simplicity, some relationships have been omitted, but are assumed to exist nonetheless. The five variables inside ovals on the left-hand side of Figure 1, are latent, exogenous variables; "latent" because they are not directly measured; "exogenous" because their causes, whatever they may be, are unanalyzed in this particular model. These latent, exogenous variables are assumed to be correlated, but the correlations in this model are taken as given, and thus not analyzed. The latent ability variable is considered to depend on father's occupation and education, mother's education, number of siblings, and respondent's sex. Not shown here is another arrow representing a residual disturbance term, which

represents all of the variation in ability not explained by the five independent variables. It is assumed to be statistically independent of the five exogenous variables, and is also assumed to be independent of the residual, disturbance terms attached to academic preparation, college plans, and educational attainment.

FIGURE 1 ABOUT HERE

HYPOTHESIZED EFFECTS

The model considers ability, as determined by a set of standardized tests administered in twelfth grade, to be dependent upon father's occupation, father's education, mother's education, number of siblings, and sex. The first three of these variables are expected to have positive effects on ability. These relationships may be causally spurious due to the omission from this model of measures by parental ability, but in any event are expected to be positive (see Scarr and Weinberg, 1978). The effect from the number of siblings is expected to be negative, primarily because we believe children from large families do not experience as much individualized attention from their parents as do children from smaller families. In making this statement, however, we do not dismiss the possibility of the effect existing because of either a birth-order effect, or a spurious effect due to less able parents bearing larger families. The effect of sex on ability is expected to be nearly zero; Wolfle (1980), for example, has summarized some of the literature on the relationship between sex and ability, and found most tests of ability were specifically constructed to produce a zero association between the test score and sex of the examinee.

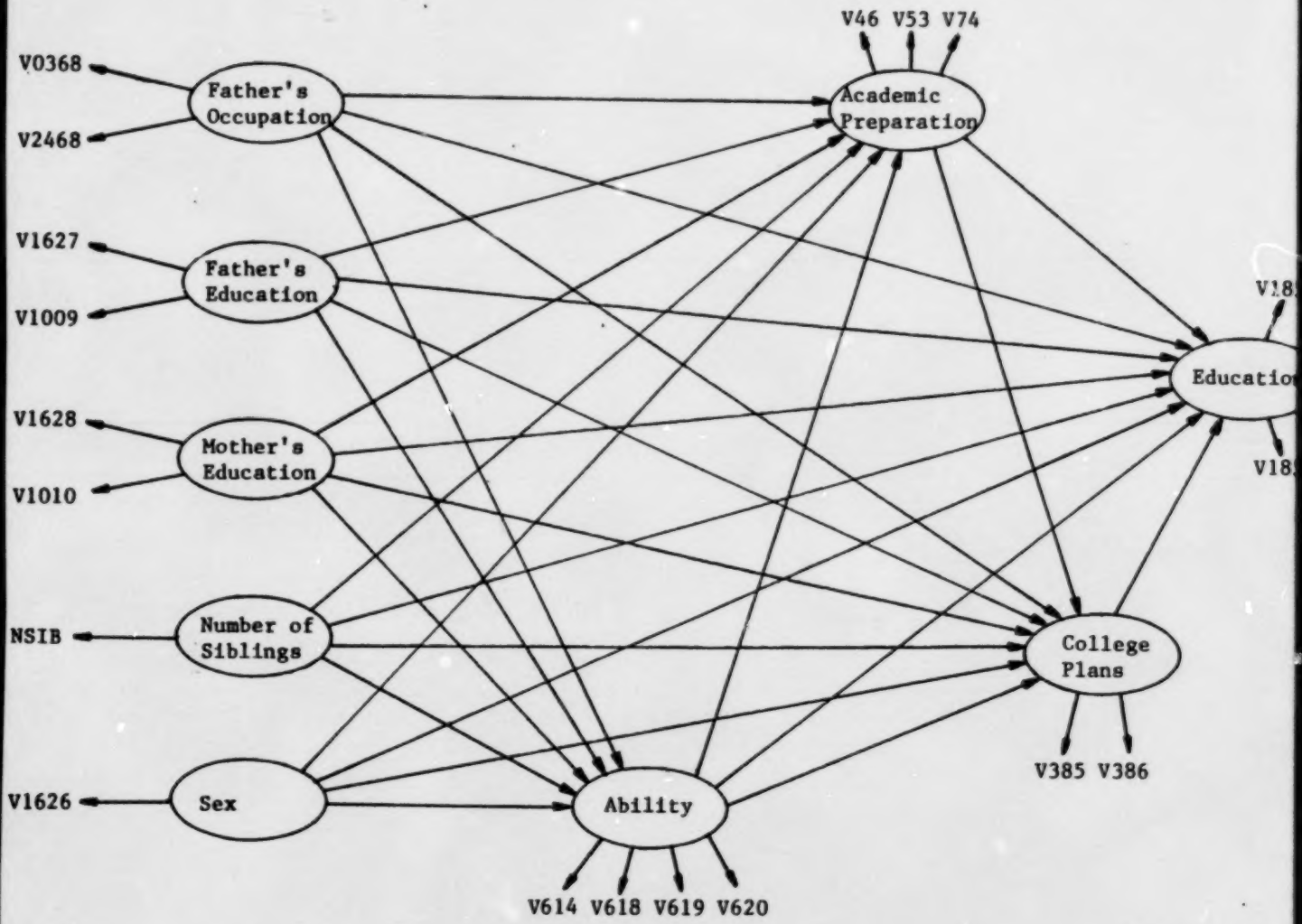

Figure 1. Structural Equation and Measurement Models of Educational Attainment Among 1972 High School Graduates

Previous studies (Heyns, 1974; Alexander and McDill, 1976), which informed our analysis, considered the influence of curriculum placement on differentiation as the major mechanism by which secondary schools function to separate students into tracks that ultimately differentiate their adult roles. Yet, as Alexander and McDill (1976) point out, it is not curriculum differentiation per se which differentiates students, but rather what happens to them in one curriculum track or another. In particular, students in college preparatory tracks complete a greater number of courses in academic subject matter, and thus develop the prerequisite skills and credentials necessary for college matriculation. In our analysis, therefore, we have eschewed the usual practice of measuring college preparatory curricular membership (a one-zero dummy variable) in favor of the number of semesters of academic courses completed in high school.

The academic preparation variable is considered to be causally determined by the five latent, exogenous variables, plus ability. We expect that students with parents of higher socioeconomic status will acquire more academic courses. We expect a negative association between academic preparation and sex, which implies women take fewer academic courses than do men. Both Heyns (1974) and Alexander and McDill (1976) report negative effects of the number of siblings on college preparatory curricular placement, and we expect to find the same in the NLS data. Finally, we expect the higher ability students to complete more academic courses than students of lower ability.

College plans measure the respondents intentions to continue their educations beyond high school. The variable may be considered dichotomously coded so that a value of unity indicates the respondent expressed plans to attend college, and a value of zero indicates no plans to attend college.

We expect positive effects from academic preparation, ability, and the three measures of parental socioeconomic status, but no a priori predictions are made about the effects of sex and the number of siblings. Despite the fact that (until very recently) more men than women actually attend college, we offer no a priori guesses about the relationship of sex to college plans, particularly when considered net of other causal influences.

Finally, we considered education to be dependent upon all of the preceding latent variables. We expect positive effects from the three measures of parental socioeconomic status. Like previous studies we expect students with more siblings to acquire less education; and we expect women to acquire less education, but we don't expect this effect to be very large in absolute value. Ability, academic preparation, and college plans will all have positive effects on educational attainment, we expect; furthermore, college plans will probably have an effect larger in magnitude than the other variables due to its immediate effect of continuing educational attainment beyond high school.

MEASUREMENT ERRORS IN MODELS OF EDUCATIONAL ATTAINMENT

The primary purpose of the present analysis is the comparison of regression slopes between whites and Mexican-Americans. We could do that with ordinary least squares regression estimates if we were willing to assume the specification bias due to measurement errors was invariant across groups. Unfortunately, this assumption is untenable. Bielby, Hauser and Featherman (1977) have shown that errors of measurement in status variables among blacks and whites are different both their magnitude and nature. They found whites to exhibit basically random errors of measurement, but blacks reported status characteristics with greater random errors, and a

nonrandum tendency to report a greater consistency among status variables than apparently was true in fact. Biases, thus, are substantially greater for blacks than for whites. As a result, ordinary least squares regression slopes when compared between blacks and whites would exaggerate racial differences as a function of differential measurement error.

Moreover, Wolfle (1979) demonstrated that Hispanics also report social background and achievement data with error, and that the magnitude of these errors exceed even those of blacks. Thus, we became convinced that comparison of coefficients across ethnic groups should be confined to coefficients corrected for differential measurement error.

The model described in Figure 1 contains latent unmeasured variables which in most cases have multiple manifest indicators. For example, the latent variable, father's occupation, has two manifest indicators, V0368 and V2468. These two manifest variables are assumed to have two causal determinants. One is from a latent, unmeasured variable, here assumed to measure "true" father's occupational status. However, the manifestly measured variable may be fallible, so a second causal effect is hypothesized. Loosely speaking we may consider this to be random measurement error, but in fact it also contains specific variation unique to each variable. Only the sum of these two components can be estimated, and while we will henceforward refer to this component as the effect of measurement error, the reader is cautioned that this is not, strictly speaking, true (see Alwin and Jackson, 1979).

The two manifest measures of father's occupation were in the first case (V0368) responses to the question administered in the base year (1972) survey which asked the respondents to indicate the kind of work done by their fathers. The categories matched, more or less, the census major occupation

groups. For this analysis the variable was recoded to the average Duncan (1961) socioeconomic index (SEI) score for the category, as revised by Hauser and Featherman (1977) to match the 1970 census occupation classification. The variable, V2468, was a composite of the individual's responses to base-year and first year follow-up questionnaire items indicating the father's occupation. This variable was coded with the revised Duncan SEI score for detailed occupation groups. The exact construction of the composite variable is given in Levinsohn, et al. (1978, pp. 76-79).

Father's education was measured with two manifest variables. The first of these was the educational composite, V1627. The second was the education question in the first follow-up, V1009. Variable V1627 was coded in such a way to represent aggregates of the categories used in V1009. To make the variables equivalent in their codes, V1009 was recoded to equal the aggregate categories of V1627. Having done that, we recoded these category codes to years of educational attainment using midpoint interval estimates taken from the U. S. Bureau of the Census (1973). These two recoded variables now represent years of father's educational attainment in which gross categories have been coded to the midpoint obtained from the distribution of educational attainment among males 25 years of age or over in 1970.

Mother's education was measured in a similar fashion. First, V1010 was recoded to the same values used in the composite V1628. These were then recoded to years of educational attainment, using the midpoint obtained from the distribution of educational attainment among females 25 years of age or over in 1970.

The number of siblings has but a single indicator, computed as the sum of variables V1460, V1461, V1462, and V1463. These questionnaire items asked the respondent to indicate the number of older brothers, younger brothers,

older sisters, and younger sisters, respectively. In handling missing data due to item nonresponse for these variables, we assumed a nonzero response to any of the four items accompanied by nonresponse to one or more of the other items represented a zero response to the nonresponse items. For example, an individual who indicated he had one older brother, but did not answer the other three questions, was assumed to have one sibling. If none of the four items was answered, the number of siblings was assumed to be missing data.

The respondent's sex was measured by the composite variable, V1626. The variable was coded 1 if male, 2 if female. As a result, positive regression estimates emanating from this variable indicate a greater value of the dependent variable for females.

The latent variable of ability was measured by four manifest indicators of achievement: V0614, a scaled vocabulary score; V0618, a scaled reading score; V0619, a scaled letter group score; and V0620, a scaled mathematics score. These variables were not recoded. They represent scores on a standardized test administered to the respondents during the spring of 1972, their senior year of high school.

The latent variable of academic preparation was measured by three manifest indicators. These were the number of semesters of science taken between July 1, 1969, and graduation (V0046), the semesters of math (V0074), and the semesters of foreign languages (V0053). Gilmartin, et al. (1976) have shown that young men who plan scientific careers completed more math and foreign language courses in high school than did young men who did not plan such careers. Moreover, young women planning such careers took more foreign language courses than would be predicted from their abilities. We expect that completion of such courses will not only predict successful attainment of plans to enter scientific careers, but will also predict the

attainment of additional years of schooling. The manifest variables were not recoded.

Two manifest variables were used to index college plans. The first of these was an NLS routing question (V0385), and was recoded unity for those people who planned to enter either a four-year college or university, or a two-year academic junior college, either full time or part time while working, apprenticing, or homemaking. Otherwise, the variable was coded zero. The second manifest measure of college plans (V0386) was based on responses to a question about what the respondents planned to do during the year after high school "if there were no obstacles". The variable was coded unity for those who said their plans were to attend either a two-year academic junior college or a four-year college or university. Otherwise, the variable was coded zero.

Educational attainment was measured with two manifest indicators. The first of these was V1854, actual educational attainment measured four years after high school graduation, and the second was V1855, planned educational attainment measured at the same time. The latent variable, educational attainment is therefore a construction not only of actual years of education completed by the third follow-up survey, but also includes a component that measures the years of additional planned education. Both manifest measures were recoded to reflect years of schooling completed or expected to be completed. Following Featherman and Carter (1976), we equated two years of attendance in a vocational, trade or business school with one year of attendance in an academic school.

SPECIFICATION OF THE LISREL MODEL

Obtaining estimates for the model shown in Figure 1 was accomplished by using LISREL (linear structural relationships by the method of maximum

likelihood), a computer program developed by Jöreskog and Sörbom (1978). The LISREL model assumes a causal structure among a set of unmeasured, latent variables, some designated as exogenous and others as endogenous. These unmeasured variables are also related to a set of observed variables such that the latent variables appear as causes of the observed variables. The LISREL model, therefore, consists of two parts: the measurement model, and the structural equation model.

These two parts of the model have been described above in the vernacular. In LISREL terminology, two random vectors, $\underset{\sim}{\eta}' = (\eta_1, \eta_2, \eta_3, \eta_4)$, and $\underset{\sim}{\xi}' = (\xi_1, \xi_2 \ \xi_3, \xi_4, \xi_5)$, represent the latent endogenous and latent exogenous variables, respectively. So that η_1 = ability, η_2 = academic preparation, η_3 = college plans, and η_4 = educational attainment. Furthermore, ξ_1 = father's occupation, ξ_2 = father's education, ξ_3 = mother's education, ξ_4 = number of siblings, and ξ_5 = sex.

The model specifies a fully recursive causal structure among the latent variables, such that:

$$\underset{\sim}{\beta}\ \underset{\sim}{\eta} = \underset{\sim}{\Gamma}\ \underset{\sim}{\xi} + \underset{\sim}{\zeta}$$

where $\underset{\sim}{\beta}$ (4x4) and $\underset{\sim}{\Gamma}$ (4x5) are matrices of structural coefficients in which $\underset{\sim}{\Gamma}$ is a full matrix relating the exogenous vector to each of the endogenous latent variables, and $\underset{\sim}{\beta}$ is a matrix relating each endogenous variable to those that follow it in the causal scheme. $\underset{\sim}{\zeta}' = (\zeta_1, \zeta_2, \zeta_3, \zeta_4)$ is a vector of randomly distributed residuals uncorrelated with each other and with $\underset{\sim}{\xi}$.

The vectors $\underset{\sim}{\eta}$ and $\underset{\sim}{\xi}$ are not observed, but $\underset{\sim}{y}' = (y_1, y_2, y_3, y_4, y_5, y_6, y_7, y_8, y_9, y_{10}, y_{11})$ and $\underset{\sim}{x}' = (x_1, x_2, x_3, x_4, x_5, x_6, x_7, x_8)$ are observed, such that:

$$\underset{\sim}{y} = \underset{\sim}{\Lambda}_y\ \underset{\sim}{\eta} + \underset{\sim}{\varepsilon}$$

and

$$\underset{\sim}{x} = \underset{\sim}{\Lambda}_x \underset{\sim}{\xi} + \underset{\sim}{\delta}$$

where $\underset{\sim}{\varepsilon}$ and $\underset{\sim}{\delta}$ are vectors of errors of measurement in $\underset{\sim}{y}$ and $\underset{\sim}{x}$, respectively. These errors of measurement represent both specific and random components of variation (Alwin and Jackson, 1979). They are assumed to be uncorrelated with $\underset{\sim}{\eta}$, $\underset{\sim}{\xi}$, and $\underset{\sim}{\zeta}$, but may be correlated among themselves. The matrices $\underset{\sim}{\Lambda}_y$ (11x4) and $\underset{\sim}{\Lambda}_x$ (8x5) are regression matrices of $\underset{\sim}{y}$ on $\underset{\sim}{\eta}$ and of $\underset{\sim}{x}$ on $\underset{\sim}{\xi}$, respectively.

Let $\underset{\sim}{\Phi}$ (5x5) be the covariance matrix of $\underset{\sim}{\xi}$. Let $\underset{\sim}{\Psi}$ (4x4) = diag (ψ_{11}, ψ_{22}, ψ_{33}, ψ_{44}) be the variance matrix of $\underset{\sim}{\zeta}$. Let $\underset{\sim}{\Theta}_\varepsilon$ and $\underset{\sim}{\Theta}_\delta$ be the covariance matrices of $\underset{\sim}{\varepsilon}$ and $\underset{\sim}{\delta}$, respectively. In application, some of the elements of the four regression matrices, and the four covariance matrices, are fixed and equal to preassigned values (often zero or unity). Other elements are free parameters to be estimated by the method of maximum likelihood. This estimation procedure requires that the estimates be maximized with respect to some known distribution, which in the case of LISREL is assumed to be the multivariate normal.

The structural model, though incomplete diagramatically, is shown in the path diagram in Figure 1. The variables enclosed in ellipses are unobserved, latent variables. The manifest variables are represented by their variable names given in Levinsohn, et al. (1978). For the sake of simplicity, in the following notations, consider X_1 = V0368, X_2 = V2468,..., X_8 = V1626, and Y_1 = V614, Y_2 = V618,..., Y_{11} = V1855. Furthermore, let $x_1 = (X_1 - \overline{X}_1)$, . . . , $y_{11} = (Y_{11} - \overline{Y}_{11})$, so that all the variables are expressed as deviations from their respective means. This transformation has no effect upon the regression slopes, but does serve to eliminate constant terms from the equations.

With this notation, the structural portion of Figure 1 is a fully

recursive model among the latent variables, represented by the following structural equations:

$$\eta_1 = \gamma_{11}\xi_1 + \gamma_{12}\xi_2 + \gamma_{13}\xi_3 + \gamma_{14}\xi_4 + \gamma_{15}\xi_5 + \zeta_1$$

$$\eta_2 = \gamma_{21}\xi_1 + \gamma_{22}\xi_2 + \gamma_{23}\xi_3 + \gamma_{24}\xi_4 + \gamma_{25}\xi_5 + \beta_{21}\eta_1 + \zeta_2$$

$$\eta_3 = \gamma_{31}\xi_1 + \gamma_{32}\xi_2 + \gamma_{33}\xi_3 + \gamma_{34}\xi_4 + \gamma_{35}\xi_5 + \beta_{31}\eta_1 + \beta_{32}\eta_2 + \zeta_3$$

$$\eta_4 = \gamma_{41}\xi_1 + \gamma_{42}\xi_2 + \gamma_{43}\xi_3 + \gamma_{44}\xi_4 + \gamma_{45}\xi_5 + \beta_{41}\eta_1 + \beta_{42}\eta_2 + \beta_{43}\eta_3 + \zeta_4$$

In algebraic form, the measurement portion of Figure 1 is:

$$x_1 = \lambda_{11}\xi_1 + \delta_1$$

$$x_2 = \lambda_{21}\xi_1 + \delta_2$$

$$x_3 = \lambda_{32}\xi_2 + \delta_3$$

$$x_4 = \lambda_{42}\xi_2 + \delta_4$$

$$x_5 = \lambda_{53}\xi_3 + \delta_5$$

$$x_6 = \lambda_{63}\xi_3 + \delta_6$$

$$x_7 = \xi_4$$

$$x_8 = \xi_5$$

$$y_1 = \lambda_{11}\eta_1 + \varepsilon_1$$

$$y_2 = \lambda_{21}\eta_1 + \varepsilon_2$$

$$y_3 = \lambda_{31}\eta_1 + \varepsilon_3$$

$$y_4 = \lambda_{41}\eta_1 + \varepsilon_4$$

$$y_5 = \lambda_{52}\eta_2 + \varepsilon_5$$

$$y_6 = \lambda_{62}\eta_2 + \varepsilon_6$$

$$y_7 = \lambda_{72}\eta_2 + \varepsilon_7$$

$$y_8 = \lambda_{83}\eta_3 + \varepsilon_8$$

$$y_9 = \lambda_{93}\eta_3 + \varepsilon_9$$

$$y_{10} = \lambda_{10,4}\eta_4 + \varepsilon_{10}$$

$$y_{11} = \lambda_{11,4}\eta_4 + \varepsilon_{11}$$

A metric for the latent variables is established by fixing some elements in the $\underset{\sim}{\Lambda}_y$ and $\underset{\sim}{\Lambda}_x$ matrices to unity. Namely, $\lambda_{41} = \lambda_{72} = \lambda_{83} = \lambda_{10,4} = 1.0$ in $\underset{\sim}{\Lambda}_y$ and $\lambda_{21} = \lambda_{32} = \lambda_{53} = 1.0$ in $\underset{\sim}{\Lambda}_x$. In $\underset{\sim}{\Lambda}_x$, λ_{74} and λ_{85} have been specified to be unity to set these latent variables exactly equal to their respective manifest indicators. As a result of these specifications, the metric of ability is measured in terms of V0620, the metric of academic preparation in terms of V0074, the metric of college plans in terms of V0385 as recoded, and the metric of education in terms of V1854 as recoded. Among the exogenous variables, the metric of father's occupation is measured in terms of V2468, father's education in terms of V1627, and mother's education in terms of V1628.

EDUCATIONAL ATTAINMENT AMONG WHITES

Assuming the joint distribution of the 19 variables in our model of educational attainment is multivariate normal, we obtain maximum likelihood estimates of parameters of the 23 structural and measurement model equations using Jöreskog and Sörbom's (1978) LISREL program. (In particular, the LISREL program we used to estimate our models employed the corrected procedures in the computation of standardized solution matrices GAMMA and D, and in t-values for multiple group comparisons.) The estimates were computed from pairwise present correlations for white 1972 high school seniors. The correlations, mean, and standard deviations among the 19 variables are shown in Table 1.

TABLE 1 ABOUT HERE

Table 1. Correlations, Means, and Standard Deviations among Variables in a Model of Educational Attainment: White 1972 High School Graduates (N = 11,743)

	V614	V618	V619	V620	V46	V53	V74	V385	V386	V1854	V1855	V368	V2468	V1627	V1009	V1628	V1010	NSIB	V1626
V614	---																		
V618	.642	---																	
V619	.412	.517	---																
V620	.525	.615	.615	---															
V46	.285	.295	.244	.417	---														
V53	.394	.353	.306	.372	.333	----													
V74	.269	.290	.268	.499	.552	.415	---												
V385	.367	.379	.315	.443	.332	.365	.351	---											
V386	.308	.314	.265	.368	.290	.296	.318	.645	---										
V1854	.407	.399	.332	.483	.352	.400	.377	.630	.525	---									
V1855	.425	.422	.318	.468	.346	.382	.366	.581	.481	.774	---								
V368	.219	.186	.154	.218	.131	.209	.169	.280	.198	.288	.283	---							
V2468	.210	.193	.165	.220	.132	.225	.174	.263	.190	.272	.289	.637	---						
V1627	.271	.249	.202	.256	.167	.251	.198	.320	.225	.355	.352	.579	.542	---					
V1009	.277	.258	.204	.258	.174	.253	.196	.326	.230	.355	.358	.574	.553	.900	---				
V1628	.253	.230	.185	.245	.162	.218	.181	.285	.196	.322	.315	.366	.349	.501	.486	---			
V1010	.264	.235	.195	.254	.164	.217	.181	.298	.213	.334	.326	.378	.357	.491	.518	.871	---		
NSIB	-.135	-.091	-.026	-.069	-.063	-.112	-.088	-.132	-.110	-.152	-.134	-.070	-.067	-.069	-.063	-.090	-.093	---	
V1626	.034	.032	.139	-.123	-.165	.112	-.196	-.029	-.018	-.929	-.081	-.035	-.013	-.017	-.028	-.047	-.047	.020	---
Mean	52.34	52.14	52.20	52.21	3.65	2.77	3.95	.480	.377	13.47	14.94	42.61	43.74	12.03	12.02	11.88	11.86	2.93	1.49
S.D.	9.63	9.28	8.76	9.34	1.91	2.27	2.00	.500	.485	1.56	2.50	22.31	22.81	3.39	3.36	2.72	2.63	1.99	.50

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When the structural and measurement models were estimated for these data, a chi-square goodness-of-fit statistic was calculated, and is shown in the first row of Table 2. This value of 5975.59 suggests at first glance that the fit of the model is not acceptable. It is well known, however, that, "in large samples virtually any model tends to be rejected as inadequate" (Bentler and Bonett, 1980). Thus, no other theoretical model will fit short of saturation, but an alternative model which is merely a specialized version of the original model, can be constructed; having estimated the two competing models, a chi-square difference test can be used to evaluate the statistical significance of the parameters that differentiate between the two competing models.

TABLE 2 ABOUT HERE

The question then becomes which parameters of our model shall be changed. First of all, the structural equation model could be altered, but it is already fully recursive and otherwise to change it would defy the logic of the temporal and theoretical relationships among these variables. Second, the factor analytic structure could be altered by allowing, for example, a causal effect of the latent factor, mother's education, to influence one or more of the manifest variables measuring father's education. But this suggestion is rejected also, again on the grounds of logic of our measurement model. Third, the initial assumption that the measurement errors among the manifest variables were uncorrelated could be relaxed. In this case, there is no objection to altering the model, because the initial assumption of zero covariance is not only restrictive, but is also contrary to the findings of Bielby, Hauser and Featherman (1977) and Wolfle (1979).

Table 2. Goodness of Fit for the Different Models of Educational Attainment of White 1972 High School Graduates (Pairwise Deletion N = 11753)

Model	χ^2	d.f.	Prob.	χ^2 for δ or ε = 0 (d.f.=1)	Prob.
Uncorrelated errors	5975.59	121	0.0		
$\delta_{5,3}$ free	5595.16	120	0.0	380.43	0.0
$\delta_{5,3}$, $\varepsilon_{10,8}$ free	5584.22	119	0.0	10.94	.001

Accordingly, we examined the off-diagonal elements of the two error variance-covariance matrices to see which elements were most likely to be different from zero. Since we wanted to find which one of the a priori assumptions of zero covariance was least probable, we relaxed the zero-restriction for that element of Θ_{ε} or Θ_{δ} which gave the largest decrease in the chi-square goodness-of-fit statistic. Following Sörbom (1975), an inspection of the LISREL-produced table of first-order derivatives (not shown here) suggested that $\Theta_{\delta_{53}}$ was probably not zero. This is the covariance between the errors of V1627 and V1628, the composite education variables for father and mother, respectively.

Allowing the covariance between δ_3 and δ_5 to be free (i.e., a parameter to be estimated within the model), the model was re-estimated, and a new chi-square goodness-of-fit statistic calculated. As can be seen in the second row of Table 2, the difference in fits between the model assuming uncorrelated errors and a new model assuming $\Theta_{\delta_{53}}$ to be nonzero was 380.43. This is distributed as chi-square with one degree of freedom, and is, of course, highly significant. Apparently in the construction of the two composite education indices, a systematic component of error was introduced into the two measures. The correlation between these two error terms was .26.

The model with $\Theta_{\delta_{53}}$ free had a chi-square goodness-of-fit of 5595.16 with 120 degrees of freedom. This is not a very good fit either; a new inspection of the derivative tables suggested $\Theta_{\varepsilon_{10,8}}$ may be nonzero. A new model was therefore estimated allowing this parameter to be free, which resulted in a chi-square of 5584.22 with 119 degrees of freedom. This goodness-of-fit is not very different from that of the previous model; as a result we adopted as our final model for whites the one as previously specified with the additional provision that the covariance between δ_3 and δ_5 is nonzero.

Measurement Model: Whites

Having found the best-fitting, most logically plausible model for whites, let us turn our attention to the measurement properties of these variables among white respondents.

As we have seen, the evidence nearly completely indicates that reporting errors for whites are random. Only one nonzero correlation among the error terms was found, and that was between two NCES-constructed composite measures of parents' education. We consider this fact significant--the correlation was not found between respondent's reports, but rather between two constructed indicators. Thus, the nonzero correlation is more likely due to some specific covariance introduced into the composites, rather than to nonrandom errors in the original reports. This general finding agrees with the results published in Bielby, Hauser and Featherman (1977), who concluded reporting errors were random for nonblack men.

Reporting errors may be random, but they are nonetheless substantial. Parameter estimates for the final measurement model for whites appear in columns 3 - 5 of Table 3. Column 6 shows the estimated reliability coefficients (the squared true score - observed score correlations estimated from the measurement model). These coefficients are striking in several ways. First, they are considerably lower than those previously reported for nonblack adults (Bielby, Hauser and Featherman, 1977) and for white twelfth-grade youths (Mason, et al., 1976). For example, Bielby, et al. (1977, p. 1258) found reliability coefficients for father's occupation of .85 and .89, while Mason, et al. (1976, p. 466) reported a coefficient of son's report of .91. In comparison, the reliability coefficients for our manifest indicators of father's occupation were only .67 and .60. We cannot explain these differences here. They may be due to differences

in the wording of the questionnaire items (see Featherman and Hauser, 1978; Kerckhoff, 1974), or to differences in personal interviews versus mailed questionnaires, or to differential coding errors, or to differential errors introduced during keypunching, or even to errors of reporting. In any event reporting errors in the NLS data seem to be considerably more severe than among other data sets which address the achievement process.

TABLE 3 ABOUT HERE

Second, it is also striking that the reliability coefficients for the background variables are greater in value for the original questions than for the NCES-constructed composites. Apparently the composites contain sources of error (additional coding? keypunching?) that are not contained in the original questions.

Third, unlike Bielby, Hauser and Featherman (1977) who found that social background variables were reported with near equal reliability, we find that parent's education was measured with considerably greater reliability than father's occupation.

Among the schooling process variables, we also find relatively low coefficients of reliability. For the four measures of ability, the reading and math tests are more reliable than the vocabulary and letter-group tests. In measuring academic preparation, semesters of math and science are considerably more reliable indicators than semesters of foreign language. In measuring college plans, a straightforward question (V0385) about college plans seems somewhat more reliable than a question (V0386) which asks respondents to consider possible obstacles to their plans.

Table 3. Measurement Model Parameter Estimates for White 1972 High School Graduates (Pairwise N=11753)

Variable		Observed Variance σ_i^2	Error Variance σ_e^2	True Variance σ_τ^2	Relative Slope λ_{ij}	Reliability Coefficient $(\sigma_\tau^2/\sigma_i^2)\lambda_{ij}^2$
True	Observed					
Ability	614	92.80	47.20	59.30	.887	.50
	618	86.10	33.83		.939	.61
	619	76.78	41.85		.768	.46
	620	87.30	28.00		1.0*	.68
Academic	46	3.65	1.91	2.36	.856	.47
Preparation	53	5.14	3.50		.836	.32
	74	4.01	1.64		1.0*	.59
College	385	.250	.056	.193	1.0*	.77
Plans	386	.235	.109		.808	.54
Education	1854	2.44	.443	1.99	1.0*	.82
	1855	6.26	1.620		1.525	.74
Father's	368	497.89	162.06	314.21	1.034	.67
Occupation	2468	520.39	206.18		1.0*	.60
Father's	1627	11.47	1.36	10.13	1.0*	.88
Education	1009	11.29	.91		1.012	.92
Mother's	1628	7.41	1.33	6.12	1.0*	.83
Education	1010	6.90	.55		1.018	.92

* Fixed value

Finally, two manifest indicators were used to measure educational attainment. One question (V1854) asked respondents in October 1976 -- four years after high school graduation -- to indicate their actual educational attainment. The second question (V1855), presented at the same time, asked respondents to indicate their educational plans. Both variables were included for the purpose of capturing both actual attainments for those who had already terminated their education, and for those who were still in the educational process. Of these two variables, actual education (V1854) was slightly more reliable than educational plans. The reliability coefficient of .82 for educational attainment is comparable to reliability coefficients from the 1973 OCG. Bielby, Hauser and Featherman (1977, p. 1258) report a reliability coefficient of .89 for the initial Current Population Survey, and .96 for the OCG reinterview (mostly by telephone). However, an OCG mailout, mailback questionnaire produced a coefficient for education of .70. Thus, the NLS reliability coefficient of .82 is greater than the comparable coefficient derived from the OCG mailed questionnaire, but less than those obtained from the OCG personal interviews.

Structural Model: Whites

In this section we present the results of the structural equation portion of our model for whites. The structural coefficients are shown in Table 4. Later we will compare these results to those obtained for Mexican-Americans; here we are concerned only with the white portion of the NLS sample. As hypothesized above, father's occupation and both parent's levels of education positively influence respondent's level of ability. These associations may be spuriously due to parental intelligence, so we do not wish to interpret these effects necessarily as the result of environmental differences in households with different socioeconomic characteristics

(see Scarr and Weinberg, 1978). The number of siblings, as hypothesized, has a negative influence on ability. Knowing one's sex does not increase one's capacity to predict one's level of ability; even with a very large number of cases, this parameter estimate is statistically nonsignificant.

TABLE 4 ABOUT HERE

As discussed above, we expected positive effects of parental status measures on academic preparation, and a negative effect from number of siblings; we thought women would complete fewer academic courses than men; we also thought higher ability students would complete more academic courses than their less able peers. In several instances, we were shown to be wrong in our assumptions. While father's occupational status had a positive influence on the number of academic courses completed, the effects from both father's and mother's education were statistically indistinguishable from zero. Women students completed about one-half fewer semesters of academic preparation than their male counterparts. The higher ability students completed more academic classes; this was the strongest of the six predictors.

Among the causal forces that influence the development of plans to attend college, Table 4 shows that ability and academic preparation are the strongest predictors. Among the background variables, father's occupation and both parent's education variables have positive and significant effects. Respondents with more siblings are less likely to express college plans. Finally, women tend to be more likely to express college plans than do men, even after controlling for all of the other independent variables in the structural equation; this effect is small, but nonetheless statistically significant.

Table 4. Maximum-Likelihood Estimates of Parameters of the Educational Attainment Process for White 1972 High School Graduates (N=11743)

Dependent Variable	Independent Variables								
	Father's Occup.	Father's Educ.	Mother's Educ.	Number Siblings	Sex	Ability	Academic Preparation	College Plans	R^2
	Standardized Coefficients								
Ability	.108*	.157*	.188*	-.075*	.012				
Acad. Prep.	.073*	.024	.020	-.050*	-.172*	.608*			
College Plans	.105*	.054*	.071*	-.068*	.035*	.292*	.305*		
Education	.017	.072*	.063*	-.040*	-.011	.164*	.124*	.565*	
	Regression Coefficients[a]								
Ability	.047 (.008)	.381 (.042)	.572 (.038)	-.262 (.037)	.185 (.145)				.16
Acad. Prep.	.006 (.002)	.011 (.008)	.012 (.007)	-.034 (.007)	-.510 (.028)	.119 (.002)			.45
College Plans	.003 (.000)	.008 (.002)	.013 (.002)	-.014 (.002)	.031 (.008)	.017 (.001)	.090 (.004)		.44
Education	.001 (.001)	.032 (.006)	.035 (.005)	-.025 (.005)	-.029 (.020)	.030 (.002)	.115 (.012)	1.786 (.042)	.68

[a]Standard errors in parentheses.

*Indicates absolute size of coefficient equals or exceeds 2.57 times its standard error.

Finally, we come to examine the estimated parameters of the structural equation for educational attainment. We had expected respondents who had fathers with greater occupational status to themselves acquire more years of schooling. However, the coefficient is not statistically significant; net of other variables in the equation, father's occupation does not influence education. Both parents' educational attainment influence the acquisition of education, with nearly equal effects. Respondents with more siblings acquire less schooling, although the effect is small once the influences of other variables are controlled. Despite the fact that women are slightly more likely than men to express college plans, in the event men and women appear to acquire nearly equal amounts of schooling, net of the influence of other variables in the model. Examining the influence of the endogenous variables, one may see that the most important predictor of years of education among high school seniors is the variable which measures plans to attend college. This may seem trivially obvious, but it is certainly not trivial in its impact. Those students who expressed plans to attend college, net of ability, academic preparation, and other background influences, will acquire 1.8 years more schooling than their peers without college plans. Completing academic courses in high school also influences educational attainment, and influences years of schooling about as much as does respondent's ability.

EDUCATIONAL ATTAINMENT AMONG MEXICAN-AMERICANS

In this section we present the results of the measurement and structural models of educational attainment among Mexican-American NLS respondents. As before, we assume that the 19 manifest variables have a joint multivariate normal distribution, and acquire LISREL estimates of the parameters using pairwise present correlations. These correlations, means, and standard deviations for Mexican-Americans are shown in Table 5. Comparison of these means to those of whites suggests that Mexican-Americans have lower levels

of socioeconomic background, and have ability test scores about 10 points less than those of whites. Mexican-American youths completed fewer semesters of academic courses than whites, and fewer Mexican-Americans expressed plans to attend college. In 1976, when asked about their planned levels of educational attainment, Mexican-Americans responded on the average that they planned 14.4 years of schooling; the corresponding figure for whites was 14.9; at the same time, Mexican-Americans had actually attained 13.0 years of schooling, while whites had attained 13.5 years.

TABLE 5 ABOUT HERE

When the structural and measurement model was estimated for the Mexican-American subsample, as shown in Table 6 a chi-square goodness-of-fit statistic was obtained equal to 440.60 with 121 degrees of freedom. Examination of the first-order derivatives suggested, as for whites, that the error terms for V1627 and V1628 may be correlated. A new model with this free parameter resulted in a chi-square of 411.91 with 120 degrees of freedom; this is a significant improvement in fit. Once again, a new examination of the first-order derivatives suggests ε_{10} and ε_{8} may be correlated. The resulting improvement of fit is marginal, even if statistically significant; because these variables are substantively unrelated, and the improvement in fit was not large in relative terms, we accepted as our final model one in which δ_3 and δ_5 are correlated. This happens to be the identical measurement model as used for whites, and suggests that neither whites nor Mexican-Americans report these variables with any substantial systematically correlated errors.

Table 5. Correlations, Means, and Standard Deviations among Variables in a Model of Educational Attainment: Mexican-American 1972 High School Graduates (N 493)

	V614	V618	V619	V620	V46	V53	V74	V385	V386	V1854	V1855	V368	V2468	V1627	V1009	V1628	V1010	NSIB	V1626
V614	---																		
V618	.600	---																	
V619	.338	.511	---																
V620	.379	.624	.566	---															
V46	.145	.203	.140	.290	---														
V53	.155	.267	.228	.270	.176	---													
V74	.064	.217	.175	.402	.526	.292	---												
V385	.271	.346	.314	.334	.181	.170	.225	---											
V386	.154	.290	.219	.333	.238	.106	.243	.615	---										
V1854	.254	.379	.247	.400	.155	.150	.259	.494	.381	---									
V1855	.256	.407	.221	.422	.262	.211	.333	.512	.463	.678	---								
V368	.096	.195	.109	.185	.062	-.092	.113	.075	.007	.094	.136	---							
V2468	.080	.113	.111	.122	-.017	-.103	.061	.060	.072	.022	.051	.634	---						
V1627	.110	.109	.007	.081	.035	.020	.129	-.013	.083	.105	.130	.508	.482	---					
V1009	.134	.166	.031	.120	.015	.050	.089	-.071	.009	.056	.101	.488	.464	.801	---				
V1628	.106	.115	-.040	.032	.017	.083	.074	.024	.066	.084	.105	.311	.280	.506	.457	---			
V1010	.123	.120	-.005	.094	-.015	.105	.076	-.010	.022	.102	.084	.352	.303	.436	.481	.802	---		
NSIB	-.259	-.154	-.070	-.146	-.082	-.008	-.038	-.058	-.054	-.067	-.105	-.143	-.192	-.114	-.116	-.156	-.159	---	
V1626	.065	-.036	.125	-.120	-.158	.075	-.184	-.088	-.084	-.021	-.079	.072	-.006	-.049	-.060	.038	.065	.061	---
Mean	44.19	43.83	45.13	44.42	2.92	2.32	3.32	.367	.315	12.97	14.39	28.84	28.47	9.35	9.48	9.11	9.18	4.51	1.49
S. D.	7.62	9.34	10.12	9.12	1.56	1.95	1.77	.482	.465	1.32	2.25	20.85	20.55	2.80	2.82	2.32	2.35	2.48	.50

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TABLE 6 ABOUT HERE

Measurement Model: Mexican-Americans

We conclude as we did for whites that reporting errors for Mexican-Americans are basically random; the only two variables with substantial covariation between their errors are NCES-constructed variables, and the systematic component of covariance could have been introduced in their construction. While the errors of measurement may be random, they are nonetheless substantial. The coefficients are shown in column 6 of Table 7. Compared to the estimated reliabilities among whites, the Mexican-American respondents report their father's occupation as reliably, or as unreliably, as among whites; reliabilities of parental education were moderately less for Mexican-Americans than for whites. As for whites, the reading and math subtests of ability were more reliable indicators than the vocabulary and letter group subtests. For academic preparation, math and science courses are more reliable than foreign language. Between the two manifest measures of education, planned education is apparently somewhat more reliable as an indicator of educational attainment for Mexican-Americans than is actual education as of October 1, 1976; this contrasts to the results for whites, for whom actual education was a more reliable indicator. These differences notwithstanding, the similarities between the groups seem noteworthy. The level of random errors seem approximately the same, and there is little evidence to suggest any extensive nonrandom reporting errors.

TABLE 7 ABOUT HERE

Table 6. Goodness of Fit for the Different Models of Educational Attainment of Mexican-American 1972 High School Graduates (Pairwise N=493)

Model	χ^2	d.f.	Prob.	$\Delta\chi^2$	d.f.	Prob.
Uncorrelated errors	440.60	121	0.0			
$\delta_{5,3}$ free	411.91	120	0.0	28.69	1	0.0
$\delta_{5,3}$, $\varepsilon_{10,8}$ free	406.43	119	0.0	5.48	1	.019

Table 7. Measurement Model Parameter Estimates for Mexican-American 1972 High School Graduates (Pairwise N=493)

Variable		Observed Variance σ_i^2	Error Variance σ_e^2	True Variance σ_τ^2	Relative Slope λ_{ij}	Reliability Coefficient $(\sigma_\tau^2/\sigma_i^2)\lambda_{ij}^2$
True	Observed					
Ability	614	58.01	36.33	49.23	.664	.37
	618	87.15	25.90		1.115	.70
	619	102.13	61.05		.913	.40
	620	83.00	33.78		1.0*	.59
Academic Preparation	46	2.43	1.47	2.20	.662	.40
	53	3.79	3.29		.478	.13
	74	3.13	.93		1.0*	.70
College Plans	385	.232	.058	.174	1.0*	.75
	386	.216	.106		.792	.51
Education	1854	1.73	.66	1.07	1.0*	.62
	1855	5.06	1.14		1.918	.78
Father's Education	368	433.54	124.75	260.84	1.088	.71
	2468	421.39	160.55		1.0*	.62
Father's Education	1627	7.85	1.67	6.16	1.0*	.78
	1009	7.92	1.37		1.032	.83
Mother's Education	1628	5.39	1.29	4.14	1.0*	.77
	1010	5.49	.71		1.074	.87

* Fixed value

Structural Model: Mexican-Americans

In this section we present the results of the structural equation portion of our model for Mexican-Americans. The structural coefficients are shown in Table 8. Here we will informally compare these coefficients to those of whites, and will postpone for the moment a formal test of statistical differences.

TABLE 8 ABOUT HERE

Among Mexican-Americans, the important predictors of ability are father's occupation and the number of siblings, but only the latter is statistically significant at the .01 level. Unlike whites, Mexican-American women scored lower on the ability factor than did men, but the net difference was not statistically significant. The detrimental influence of more siblings was greater among Mexican-Americans than among whites, but the influence of parental education and father's occupation among Mexican-Americans were statistically indistinguishable from zero.

In the determination of academic preparation, among Mexican-Americans none of the social background variables were statistically significant, but ability and sex were. Increments to ability produced additional academic preparation of the same order of magnitude as among whites. The Mexican-American men, like the whites, completed more academic courses than did women. Note here, as in the equation predicting ability, that only about one-half as much variance is explained for Mexican-Americans than for whites.

In developing plans to attend college, the Mexican-Americans, as do whites, depend primarily on ability and academic preparation. None of the

Table 3. Maximum-Likelihood Estimates of Parameters of the Educational Attainment Process for Mexican-American 1972 High School Graduates (N=493)

Dependent Variable	Independent Variables: Father's Occup.	Father's Educ.	Mother's Educ.	Number Siblings	Sex	Ability	Academic Preparation	College Plans	R^2
Standardized Coefficients									
Ability	.183	.007	.007	-.130*	-.023				
Acad. Prep.	-.024	.051	.052	.036	-.197*	.443*			
College Plans	.112	-.254	.057	.031	-.079	.450*	.167*		
Education	-.083	.145	.016	-.013	.031	.166*	.114*	.437*	
Regression Coefficients[a]									
Ability	.080 (.037)	.016 (.252)	.023 (.221)	-.455 (.142)	-.351 (.693)				.07
Acad. Prep.	-.002 (.008)	.024 (.053)	.031 (.047)	.025 (.030)	-.582 (.147)	.087 (.012)			.22
College Plans	.003 (.002)	-.036 (.014)	.010 (.012)	.006 (.008)	-.069 (.040)	.026 (.004)	.049 (.018)		.30
Education	-.007 (.004)	.064 (.030)	.009 (.026)	-.008 (.017)	.085 (.084)	.030 (.009)	.106 (.038)	1.381 (.171)	.58

[a]Standard errors in parentheses.

*Indicates absolute size of coefficient equals or exceeds 2.57 times its standard error.

social background variables are statistically significant. In this instance, the two groups seem to be nearly the same.

Examining the coefficients for educational attainment, we are struck more by the similarities between the two groups than by their differences. Among the background variables, there are some differences, but it cannot be said that they consistently favor or disfavor either group. Thus, for neither group is the effect of father's occupation statistically significant; for Mexican-Americans the influence of father's education is nearly twice that exhibited among whites, but the influence of mother's education among Mexican-Americans is negligible. Additional brothers or sisters is less a detriment to Mexican-Americans than to whites, but in neither case is the effect very large. For neither group are either men or women advantaged in terms of educational attainment. For both groups, the most important predictors of education were the endogenous variables, ability, academic preparation, and college plans; but the marginal rates of returns for these variables were nearly the same for the two groups. The only difference of note was for the variable college plans; whites who expressed plans to attend college actually attained 1.8 additional years on the average; Mexican-Americans attained only 1.4.

In sum, we are impressed more by the similarities of effects between Mexican-Americans and whites than we are by the differences. The coefficients do, of course, vary between the two groups, but the direction of statistically significant effects are most often identical, and the relative magnitudes are often close in value. The question that remains unanswered is whether the differences that do exist between the two groups reflect substantive differences, or whether they are differences that might have been expected to occur by chance. This is the question to which we

now turn our attention.

COMPARISON OF STRUCTURAL EFFECTS ACROSS ETHNIC GROUPS

Having obtained estimates of the parameters of the model of educational attainment for whites and Mexican-Americans, we then asked whether the differences in estimated parameters resulted from random sampling fluctuations, or whether the differences resulted from real differences in the process of educational attainment between the two groups. To effect this analysis we estimated the model for both groups, specifying the same measurement and structural models, but allowing the estimates of these parameters to vary as they will across the two groups. We then constructed a new model, specifying that the parameter estimates in the gamma matrix (effects from exogenous to endogenous variables) and beta matrix (effects from endogenous to subsequent endogenous variables) be invariant across the two groups. If the goodness-of-fit statistics between these two models do not vary significantly we would conclude that specifying invariant structural effects across the groups did not seriously erode our ability to fit the model to the data. If, however, the two chi-square values were significantly different, then we would conclude there were differences between the groups large enough to seriously erode the model's ability to reproduce the observed variance-covariance matrix. We were in fact testing for statistical interactions among the structural coefficients across the two groups. Rejecting the hypothesis of invariant gamma and beta coefficients was therefore equivalent to concluding that the structural coefficients of the process of educational attainment vary between whites and Mexican-Americans.

When both groups were considered together, and the gamma and beta matrices specified to be invariant across the groups (no differences in the process of educational attainment), a chi-square goodness-of-fit statistic

was obtained equal to 6146.12 with 267 degrees of freedom. This reflects, of course, a poor fit, but the question of interest is whether the fit is any less worse than a model that does not constrain the gamma and beta coefficients to be invariant across groups. When such a model was estimated, the chi-square coefficient was 6035.76 with 241 degrees of freedom. The difference between these values is equal to 110.36, which is also distributed as chi-square with 26 degrees of freedom. At the .01 level of probability, we conclude the chi-square value is statistically significant. That is, the coefficients in our model of educational attainment differ to a degree not attributable to sampling error.

SUMMARY AND DISCUSSION

We conclude that the process of educational attainment varies between whites and Mexican-Americans to a statistically significant degree. But now we want to draw the common distinction between differences that are statistically significant, and differences that are substantively important. It is well known that practically any difference, no matter how small, may be statistically significant if the sample size (a value which enters into the denominator of the test statistic) is large enough. In the NLS sample, the sample size is very large indeed. There were nearly 12,000 whites in the sample. We may, as a result, be very confident in the stability of parameter estimates for whites; a confidence which extends to the comparison of these coefficients to corresponding coefficients for Mexican-Americans. It is another question, however, whether these differences are substantively important. As we have noted throughout this paper, we have been struck more by the similarities between the two groups than we have been by the differences. Thus, our overall impression is that the process of educational

attainment may differ between the two groups, but not a degree worthy of much notice.

Yet we want to be clear about whom we are making these statements. The population to which these results are generalizable consists of high school seniors. Our conclusions pertain to students who were still in school in the spring of their senior year (1972): when we examine the social mechanisms by which these people convert their human capital into years of additional schooling, it does not seem to matter very much whether a person is white or Mexican-American. Those high school seniors who went on to accumulate additional years of postsecondary schooling were those who possessed higher ability scores, who accumulated academic courses in high school, and who had developed plans to attend a college or university.

It is one thing to say that the process of educational attainment varies but little between the two groups. It is another to say that the outcomes will be the same. Because Mexican-Americans have mean ability scores nearly ten points below those of whites, take fewer academic courses in high school, have fewer members who plan to attend college, have parents of lower socioeconomic status, and more siblings, thus will the 1972 high school cohort of Mexican-Americans attain less schooling than their white peers. Equality of educational opportunity will not equalize outcomes when the groups are not equal to begin with. Before Mexican-Americans can be expected to match whites in terms of educational attainment, either the groups must be equalized in terms of socioeconomic and school-processing variables, or the process by which these variables translate to years of schooling must be unequalized. Strangely enough, this means decreasing the influence of these variables upon schooling for Mexican-Americans; as long as Mexican-Americans have mean values on the independent variables less than

those of whites, structural effects equal to those of whites mean less schooling for Mexican-Americans.

Until the completion of this analysis, there was no comprehensive comparison of the process of educational attainment between majority whites and Mexican-Americans; particularly when one focuses on a recent cohort of high school graduates, and measures differences in structural coefficients net of differences in measurement error. When we did so, we found the process of educational attainment essentially invariant between whites and Mexican-Americans. The recent entry of this cohort into the labor market has precluded our extension of the analysis to the examination of the processes of achievement of occupations and earnings. From what we know about differences in these processes for whites and blacks, however, it would be premature to make any conclusions about differences between whites and Mexican-Americans in the achievement of occupations and earnings on the basis of these results.

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